Bargaining Power, Strike Durations, and Wage Outcomes: An Analysis of Strikes in the 1880s

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Strike outcomes in the 1880s had a "winner-take-all" character. Successful strikes ended with a discrete wage gain; failed strikes ended with a return to work at the prestrike wage. We present a theoretical interpretation of these outcomes based on a war-of-attrition model. We fit an empirical model specifying the capitulation times of the two parties and the size of the wage gain in the event of a strike success. The results show a systematic relation between the determinants of strike success and the determinants of the wage gain for a successful strike.

Until well into the twentieth century, the U.S. Bureau of Labor classified strikes and lockouts by their relative success. Most disputes in the 1880s and 1890s were either won or lost: a surprisingly small fraction were recorded as ending with a compromise or partial success.[1] Although the identification of winners and losers appears slightly contrived to modern observers (Kennan 1986), government statisticians and academics at the turn of the century made extensive use of the classification.[2] These analysts

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[1] See Peterson (1937). Compromises made up only 5%–15% of strike settlements in the period 1881–1900.

[2] See Adams (1905), Cross (1908), and Moore (1911), for example.

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evidently viewed the distinction between successful and failed strikes as a natural and empirically useful taxonomy.

This article presents a detailed analysis of strike outcomes in the data set collected by the Bureau of Labor for the period 1881–86. Our analysis begins with a simple empirical observation: in disputes arising over demands for a wage increase, successful strikes almost always resulted in a significant wage gain, while failed strikes almost always ended with no change in wages. Rather than an arbitrary distinction between more or less favorable outcomes, the classification of successful and failed strikes in the 1880s reflected an inherent discreteness in the nature of strike settlements.

We argue that this discreteness reflects the institutional structure of the late nineteenth-century labor market. Unlike the situation today, workers involved in strikes over wage increases in the 1880s were typically *not* members of a secure union with recognized bargaining rights. Instead, the outcome of the strike determined their collective bargaining status. In modern (post–National Labor Relations Act) terminology, we interpret strikes over wage increases in the 1880s as primarily strikes over "union recognition."[3] The main question resolved by strikes was whether the employer would recognize employees' bargaining power. If so, a discrete wage premium was established. If not, wages and working conditions returned to their preexisting levels.[4]

In the spirit of this interpretation, we present a theoretical model of strike durations and wage outcomes based on a war-of-attrition model (Maynard Smith 1974).[5] We interpret the wage premium that is potentially earned by an effective "union" (or combination of workers) as a prize captured by the winner of the strike. Following Kennan and Wilson (1989), we model the delay costs of the disputants as random variables that are asymmetrically observed. A strike continues until one of the parties concedes the prize to the other. The optimal strategies of the parties determine a pair of capitulation times (or "holdout" times) that depend on the size of the prize, the actual delay costs of each party, and each party's expectations about its rival's costs.

In this model the resolution of a strike reveals the capitulation time of the losing party and, if the strike is won by workers, the size of the wage

[3] Writing at the time, Ely (1886) argued that strikes were more prevalent in the United States than in England because the weaker American unions were constantly fighting for employer recognition (pp. 151–52).

[4] Interestingly, bargaining for a first contract under the current institutional structure has a similar "discreteness." In the 1980s, unions failed to achieve a first contract in about 30% of new certifications (see Cooke 1985). In cases where a first contract was achieved, unions typically raised wages by 5% (Freeman and Kleiner 1990).

[5] See Craig (1989) for an earlier attempt to use a war-of-attrition model to study strike data for 1881 and 1891 from New York State.

premium. Accordingly, we fit a three-equation model specifying the capitulation times of workers and employers and the size of the wage increase if workers establish effective bargaining power. The exogenous variables in our empirical specification include the number of strikers; the fraction of the firm's employees involved in the strike; the fraction of female workers in the firm's workforce; whether or not the strike was ordered by a labor organization; and controls for industry, occupation, location, and time.

The estimated effects of these variables on the probability of a successful strike are generally consistent with earlier investigations of strike outcomes in the 1880s (e.g., Friedman 1988). Strikes ordered by a labor organization, strikes with fewer female workers, strikes initiated prior to the wave of unrest following the Haymarket incident in May 1886, strikes in the building trades and the shoe industry, and strikes involving a larger fraction of the firm's workforce were more likely to succeed. Interestingly, all of these factors raise the wage conditional on a successful strike. We interpret this finding as evidence that employers with greater potential rents had higher costs during a work stoppage.

Our conclusion that strikes were more likely to succeed against employers with greater rents sheds interesting light on labor union policies in the decades after 1886. By concentrating their organizing effort on workers in occupations and industries with greater potential wage gains from unionism, union leaders in the late 1880s and 1890s may have been maximizing their survivability.

I. Historical Overview

This section presents a brief description of the 1880s labor market and the strike data collected by the Department of Labor.[6] The early 1880s marked a return to relative prosperity after the prolonged depression of the 1870s. The labor movement experienced a parallel resurgence: from a low point following the wave of unsuccessful strikes during 1877, membership in traditional craft-based labor organizations grew steadily between 1880 and 1886 (see Wolman 1924, chap. 2). Far more spectacular was the surge in membership in the Knights of Labor. Spurred by a successful railway strike in 1885, the Knights grew into a powerful national movement claiming some 700,000 members by late 1886. The Knights' admission of

[6] See Commons and associates (1926), Ware (1929), and Taft (1964) for comprehensive histories of the labor movement in the 1880s; Wolman (1924) for a study of the growth of trade union membership; Ulman (1955) for an analysis of factors leading to the rise of national trade unions during the late 1800s; David (1936) and Avrich (1984) for studies of the Haymarket affair; Groat (1905) for an analysis of the legal status of strikes in the 1880s; Sundstrom (1990) and Hanes (1992) for studies of wage determination in the late nineteenth century; Griffin (1939) for a descriptive overview of strike activity in the United States until 1930; and Friedman (1988) for a quantitative analysis of strike success in the 1880s and 1890s.

unskilled and semiskilled workers (including female and black workers) was a significant departure from earlier trade union policies and brought a new class of workers into the labor movement and onto the picket line.

A surge of labor unrest spread over the country in early 1886, culminating with the call for a "general strike" for an 8-hour working day. Backed by the Federation of Organized Trades and Labor Unions (the precursor of the American Federation of Labor) and a number of radical labor organizations, workers in many cities struck on May 1, 1886 (see David 1936). In Chicago, the ensuing demonstrations led to a violent confrontation between police and strikers at Haymarket Square on May 3. Public reaction to this incident bolstered employer opposition to union organizing, and bitter confrontations continued throughout the year.[7]

In response to the wave of "labor problems" in 1886, the Bureau of Labor attempted to enumerate every strike and lockout in the United States between 1881 and 1886. The bureau compiled a list of disputes from newspapers and trade magazines and then assigned field agents to track down details of the known disputes and gather information on other strikes or lockouts during the period. According to the commissioner, "the parties instigating a strike were consulted . . . and the agent, after considering all the evidence to be gained on either side, reported what the facts seemed to be" (U.S. Department of Labor 1888, p. 10).[8]

The results of the bureau's inquiries are tabulated in the *Third Report of the Commissioner of Labor* (U.S. Department of Labor 1888). Information is provided on about 5,000 individual disputes, including the location and industry of the employer, the number of employees affected by the dispute, average wages and hours before and after the stoppage, the cause of the dispute, the beginning and ending date of the stoppage, and the resolution of the dispute.[9]

By modern standards, the data collection effort underlying the preparation of the *Third Report* was extraordinary. Nevertheless, a recent study of the accuracy of the strike listings suggests that the Bureau of Labor did not achieve a complete census of disputes. Bailey (1991) compares the

[7] Other notable labor disputes in 1886 included a second strike between western railway companies and the Knights of Labor (see Taussig 1887; and U.S. Department of Labor 1888, pp. 29–32) and the lockout of meat packers in Chicago in the fall of 1886 to restore the 10-hour working day (Perlman 1922, pp. 97–98).

[8] We have no direct evidence on how many field agents were employed by the bureau or on what fraction of cases the agents managed to successfully interview the parties to a strike.

[9] Some of the Bureau's definitions and methods of reporting are described in the *Third Report.* Later reports describing strike activity in 1887–94 (*Tenth Report of the Commissioner of Labor*), 1894–1900 (*Sixteenth Report of the Commissioner of Labor*) and 1901–5 (*Twenty-first Report of the Commissioner of Labor*) contain further information on the Bureau's methodology and reporting methods (U.S. Department of Labor 1896, 1901, 1906).

strikes listed in the *Third Report* and the later *Tenth Report* (U.S. Department of Labor 1896) to those mentioned in local newspapers in Terre Haute, Indiana, over the period 1881–94. His results suggest that only one-half of the 46 strikes in this 15-year period were actually recorded by the Bureau of Labor. This undercount poses no particular problem for our statistical analysis, provided that the bureau enumerated a random sample of disputes. Strikes that were overlooked by the Bureau of Labor varied in duration and size, and Bailey was unable to find any strong pattern differentiating strikes that were excluded from *Reports* from those that were included. We agree with his conclusion that final judgments about the representativeness of the strikes reported in the *Third Report* will require additional research in other cities. To the extent that the uncounted strikes differ from the recorded strikes, however, the available sample may present a biased picture of strike outcomes in the late nineteenth century.

II. A Preliminary Descriptive Analysis

From the listings in the *Third Report,* we elected to analyze strikes from Illinois, New York, and Massachusetts. These three states accounted for 41% of all disputes (strikes and lockouts) in the United States in the early 1880s. During this time period, lockouts represented only 6% of total disputes, and for convenience we have excluded them from our analysis.[10]

The *Third Report* lists 2,256 individual strikes in Illinois, Massachusetts, and New York between 1881 and 1886. Initial analysis of these data suggested a series of clerical errors affecting 77 strikes in Illinois.[11] Exclusion of these strikes generates a usable sample of 2,179 observations. Table 1 presents some simple descriptive statistics for the sample, including breakdowns by state and year. Just over one-half of the strikes in the sample are drawn from New York state, while 30% are from Illinois and 15% from Massachusetts. The annual number of strikes is fairly stable from 1881–85 and then shows a dramatic increase in 1886. Much of this increase grew out of the "8-hour-day" campaign launched by the Federation of Trades and Labor Unions. The number of strikes in March and April of 1886 was over twice the average for these months in the preceding 5 years. In May 1886 there were as many strikes as in all of 1881 and 1882 combined.

Rows 2–6 of table 1 show the median number of workers involved per strike, the average fraction of workers at each establishment involved in the strike, the average daily prestrike wage of strikers, the percentage of

[10] A cursory examination suggests that lockouts were longer than strikes (mean duration 38 days vs. 20 days) and were less likely to be ordered by a labor organization (51% vs. 77%), but were about equally likely to result in a "success" for workers (49% vs. 47%).

[11] These strikes all involved laborers and wharf hands in Illinois. A list of excluded strikes is available on request.

Table 1
Mean Characteristics of Strikes, by State and Year

	All (1)	State: Illinois (2)	State: Massachusetts (3)	State: New York (4)	Starting Year: 1881 (5)	Starting Year: 1882 (6)	Starting Year: 1883 (7)	Starting Year: 1884 (8)	Starting Year: 1885 (9)	Starting Year: 1886 (10)
1. Number of strikes	2,179	705	300	1,174	219	207	205	234	299	1,015
2. Median number of strikers	50	80	35	40	60	60	75	35	60	50
3. Average fraction involved (%)	80.6	89.7	41.9	85.0	84.4	83.4	84.6	83.9	75.7	79.1
4. Average previous wage ($/day)	2.12	1.99	2.01	2.22	1.97	2.15	2.12	2.43	2.08	2.08
5. Ordered by labor organization (%)	76.8	74.0	57.7	83.3	69.4	67.6	79.0	77.8	73.6	80.5
6. Average fraction female strikers (%)	9.8	1.8	25.5	10.7	8.1	7.0	12.1	8.7	14.0	9.4
7. Involving generic employees (%)	25.6	44.3	28.7	25.6	25.1	25.1	16.6	23.5	27.8	41.4
Causes of strike:										
8. For a wage increase (%)	47.1	38.2	54.0	50.7	69.4	63.3	53.7	39.3	50.8	38.4
9. Against a wage cut (%)	11.6	12.1	15.7	10.2	10.0	11.6	18.5	24.8	18.4	5.4
10. For a change in hours (%)	18.3	30.9	3.3	14.6	5.4	4.3	1.0	11.5	1.0	34.1
Outcomes:										
11. Median duration (days)	9.0	10.0	14.0	7.0	7.0	7.0	8.0	7.0	11.0	10.0
12. Successful (%)	46.9	31.2	41.7	57.7	52.1	52.7	57.6	58.1	55.9	37.2
13. Compromise (%)	9.2	12.2	9.3	7.3	13.7	6.3	5.4	2.6	3.7	12.7
14. Strike breakers employed (%)	38.7	38.6	33.0	40.3	41.0	36.7	47.8	34.6	37.1	38.2
15. All strikers replaced (%)	6.9	.9	12.0	9.2	5.0	3.9	9.3	12.0	5.4	6.7

NOTE—Sample is drawn from *Third Annual Report of the Commissioner of Labor* (U.S. Department of Labor 1888) and includes all strikes in Illinois, Massachussetts, and New York occurring between 1881 and 1886.

strikes that are recorded as having been ordered by a labor organization, the average fraction of female employees at each establishment involved in the strike (prior to the strike), and the percentage of strikes involving an unspecified occupational group (for simplicity we refer to these as strikes involving "generic employees").[12]

The size distribution of strikes in the sample is right-skewed, with many small strikes and a few very large disputes. As a consequence, the median size of strikes (50 employees over all years and states) is perhaps more informative than the mean size (245 employees over all years and states). On average, 80% of employees at affected establishments participated in the strike.

The average wage of strikers in the 1880s was approximately $2.00 per day, although the figure ranges from under $0.75 to over $4.00 per day.[13] An analysis of prestrike wages reveals that earnings were lower for generic employees and for groups with a larger fraction of female workers and varied significantly across industries.[14] Since our statistical models (below) include year effects, we have not attempted to adjust nominal wages for the modest fall in prices between 1881 and 1886.[15]

On average, three-quarters of strikes in the 1880s were ordered by a labor organization.[16] The fraction of such "authorized" strikes is higher in New York State, lower in Massachusetts, and shows a slight upward trend during the sample period. The average fraction of female employees (row 6) is fairly constant over time but varies across states, with a relatively high fraction in Massachusetts (mainly in the textile and boot and shoe industries) and a very low fraction in Illinois. The fraction of strikes involving generic employees is higher in Illinois than New York or Massachusetts and is also higher in 1886 than in earlier years.

[12] Unspecified employee groups may involve either unskilled workers or a broad range of occupation groups (or both). We have no strong reason to believe strikes by "generic" workers were different from strikes by others. We include this variable because it was a worker classification that was thought to be relevant to contemporary observers. This variable was generally not statistically significant in the estimated models.

[13] These rates are comparable to other wage data for the period. For example, Long's (1960) tabulations of manufacturing wages for 1880 show average rates of $2.20–$2.45 per day for skilled occupations and $1.32 per day for laborers.

[14] The coefficient of the fraction female variable indicates a 40% wage gap between male and female workers, controlling for industry, location, and time effects. Generic employees earned 6% less than other employee groups.

[15] Lebergott's price index (U.S. Bureau of the Census 1975, series D737) shows constant prices between 1881 and 1882, a 4% decline in prices in 1883, a 3% decline in 1884, and then relative stability between 1885 and 1886.

[16] According to the *Twenty-first Report* (U.S. Department of Labor 1906, p. 31) strikes not ordered by a labor organization included strikes of previously unorganized workers as well as strikes initiated by members of labor organizations but without the authority of these organizations.

Rows 8–10 of table 1 give the fractions of strikes attributable to three major causes: workers' demands for a wage increase, employers' demands for a wage cut, and workers' demands for a reduction in hours. The remaining 23% of strikes are attributable to a variety of causes including employee discharges, changes in work rules, and sympathy strikes. The importance of the 8-hour-day campaign in 1886 is illustrated by the unusually high fraction of hours-related strikes in that year. The large number of hours strikes also accounts for the rise in the fraction of generic employees in 1886.

Information on strike durations and outcomes is presented in rows 11–15.[17] Strike durations are right-skewed, implying a mean duration (20 days) considerably in excess of the median (9 days). The most frequent duration is 1 day (12% of all strikes); one-third of all stoppages ended within 3 days. Close to one-half of all strikes were successful, while 40% were failures. Only a small minority of strikes ended with a compromise between the positions of workers and the employer.

Another measure of strike outcomes is the extent to which outsiders were recruited to replace the strikers. In 40% of strikes at least some outside replacements were employed at the end of the dispute, and in 7% of strikes all the strikers were replaced or the employer closed down. In most cases, however, outside replacements accounted for a relatively small fraction of poststrike employment.[18]

Table 2 provides descriptive information on strikes by the cause of the dispute. The largest single category of strikes are those over wage increases, and we concentrate on these strikes in the remainder of the article. Wage-increase strikes share similar characteristics to other disputes, although they are more evenly distributed over the sample period. Strikes against wage cuts tended to be longer than strikes for a wage increase but about equally as likely to succeed. Hours strikes tended to be less successful than other strikes and also tended to involve fewer female workers and a greater fraction of unspecified employee groups.

[17] The Bureau of Labor deemed a strike to be over when the employer was "open and operating as usual" (U.S. Department of Labor 1896, p. 15). The *Third Report* does not explicitly describe the bureau's system for distinguishing successful and failed strikes. According to the *Twenty-first Report,* however, successful strikes were those in which all the strikers' demands were granted, failed strikes were those in which none of the strikers' demands were granted, and partially successful strikes were those in which some of the strikers' demands were met (U.S. Department of Labor 1906, p. 79).

[18] In our earlier version of this article, we also estimate a model that includes whether or not the firm hired strike replacements. As we noted, this decision by the firm is unlikely to be exogenous to the capitulation times of the parties. For this reason, the results we report here exclude this variable. See Card and Olson (1992) for these estimates.

Table 2
Mean Characteristics of Strikes, By Cause of Dispute

	Strikes for Wage Increase (1)	Strikes against Wage Cut (2)	Strikes for Hours Cut (3)	Miscellaneous Causes (4)
1. Number of strikes	1,026	252	399	502
2. Median number of strikers	50	75	60	42
3. Average fraction involved (%)	78.6	80.3	89.3	77.9
4. Average previous wage ($/day)	2.01	2.06	2.27	2.24
5. Ordered by labor organization (%)	74.1	67.9	93.0	73.9
6. Average fraction female strikers (%)	11.0	16.5	1.5	10.8
7. Involving generic employees (%)	25.7	28.6	57.4	26.7
Timing:				
8. January–April 1886	17.8	6.7	6.8	15.1
9. May 1–7, 1886	8.0	2.0	57.1	4.4
10. After May 7, 1886	12.1	13.1	22.8	25.3
Outcomes:				
11. Median duration (days)	7	14	11	10
12. Successful (%)	51.9	48.0	32.6	47.4
13. Compromise (%)	11.4	6.7	13.0	2.8
14. Strike breakers employed (%)	36.9	31.7	31.3	51.8
15. All strikers replaced (%)	5.4	8.7	4.8	10.8
16. Average change in log wages (%)	8.4	−4.9	.0	.0
17. Fraction with no change in wages	31.1	47.6	78.9	74.7
18. Average change in weekly hours	−.5	−.1	−3.8	−.1
19. Fraction with no change in hours	91.5	95.6	55.9	95.4

SOURCE.—See table 1 note.

Rows 16–19 show the average changes in wages and hours associated with strikes for various causes. We measure the wage change by the difference between the poststrike and prestrike wage rate of striking employees. The hours change is a similar difference in the weekly hours of workers affected by the strike (no separate hours data are available for the strikers themselves). For strikes over wage increases, the average wage change is relatively large and positive, while for strikes against wage cuts, the average wage change is negative. In a sizeable fraction of strikes in either category, however, the stoppage ended with no change in wages.

More insight into this fact is provided by figure 1, which plots the frequency distribution of wage changes for strikes over wage increases. The distribution is bimodal, reflecting a mixture of strikes that ended with no change in wages (mostly failed strikes) and strikes that ended with a significant wage gain (mostly successful strikes). A similar pattern appears in the frequency distribution of wage changes following strikes against wage cuts. This distribution is a mixture of a "spike" at zero (reflecting the successful strikes) and a single-peaked distribution of negative wage changes (reflecting the failed strikes).

Figure 1 suggests that strikes over wage increases were usually resolved by a "winner-take-all" settlement. If the strike was successful, a strictly positive wage gain was achieved. If the strike failed, the wage returned to its prestrike level. The average wage increase conditional on a successful strike was 13.6%—roughly equal to the union wage premium in the modern labor market (Lewis 1986) and similar to Eichengreen's (1987) estimate of the union wage effect for Iowa workers in 1894.

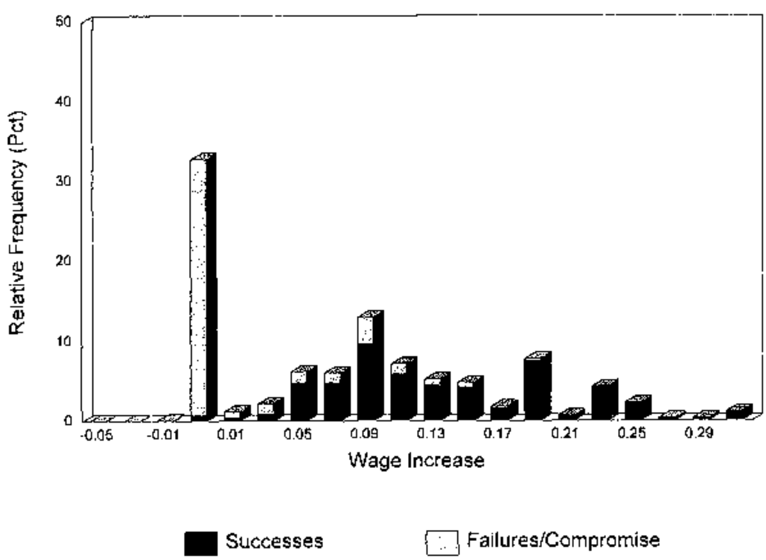

FIG. 1.—Distribution of wage increases

Further evidence on the nature of strike settlements is presented in table 3. Here we have tabulated the outcomes of wage increase strikes by six duration categories. There is a strong association between the length of the strike and the likelihood that it succeeded (col. 2) or failed (col. 3). Conditional settlement rates for the set of ongoing disputes at the beginning of an interval are shown in columns 4–6. Over the duration intervals in the table, there is a decreasing rate of strike settlements.[19] The main factor in this decreasing hazard is the declining probability of a successful settlement: from 7% per day during the first 3 days of a strike to less than 0.4% per day after a month or more.

Despite the decreasing likelihood of a successful settlement, the wage increase conditional on a success is unaffected by the duration of the strike (col. 7). The wage change conditional on a failed strike is similarly unaffected (col. 8).[20] Thus, the bifurcation of wage settlements following successful and unsuccessful strikes persists even after controlling for the duration of the dispute. The average wage increase conditional on the duration of the strike declines steadily, however, reflecting the declining fraction of successful settlements (col. 7).

III. A Theoretical Model of Strike Outcomes

Building on the descriptive evidence in the previous section, this section presents a theoretical model of strikes over wage increases in the 1880s. The model is a "war-of-attrition" model (Maynard Smith 1974; Kennan and Wilson 1989) with two possible outcomes: either the strike is won by employees, in which case a wage premium is established (for some unspecified time into the future), or the strike is won by the employer, in which case the strikers return to work at the previous wage. In this setup, the bifurcation of wage settlements following successful and failed strikes is attributed to the difference between wages in the presence or absence of an effective "union."[21]

In addition to providing a simple explanation for the contrast between successful and failed strikes, this model is consistent with a variety of qualitative evidence on the nature of labor disputes in the late nineteenth century. Many disputes were informally organized with little or no control by an extant union leadership. Even in cases where a strike was ordered by a labor organization, the fraction of employees supporting the walkout was in doubt. Ehrlich (1974) concludes that "many strikes were hampered by the disintegration of the united front put forward at the outset of the

[19] Daily settlement patterns over finer time intervals reveal "spikes" in the settlement rates at 7, 14, and 21 days—see Card and Olson (1992).

[20] Only 22 of the 322 failed strikes have a nonzero wage change.

[21] We interpret the term "union" broadly to include any collective organization of employees with control over the supply of labor to the employer.

Table 3
Strike Outcomes and Settlement Rates by Duration (Strikes over Wage Increases Only)

		Strike Outcomes		Average Daily Settlement Rates*			Average Percentage Wage Change:		
	Number of Strikes (1)	% Successful (2)	% Failed (3)	Success (4)	Failure (5)	Compromise (6)	Overall (7)	If Successful (8)	If Failed (9)
Duration:									
All	1,026	51.9	36.6	...	...	...	8.4	13.6	−.0
1–3 days	305	68.9	22.6	6.8	2.2	.8	10.8	14.9	−.1
4–7 days	224	61.6	30.4	4.8	2.4	.6	8.3	12.7	−.2
8–14 days	168	41.7	39.3	2.0	1.9	.9	7.2	12.2	.2
15–28 days	135	43.0	42.2	1.3	1.2	.4	6.4	12.3	.1
29–90 days	121	37.2	50.4	.4	.5	.1	6.3	14.5	−.4
90+ days (complete)	18	44.4	33.3	...	...	...	9.3	11.6	.0
All replaced†	55	.0	100.0	...	...	...	...	...	...

NOTE.—See table 1 for sources and definitions.
* Conditional probabilities of settlement ending with success, failure, and compromise. Probabilities expressed at daily rates.
† Strikes in which all strikers were eventually replaced. Duration and wage information are unavailable.

stoppage by the strikers. After varying lengths of time, men who had originally stood with their fellow workmen weakened and returned to work."[22] To reach a successful settlement, workers had to convince their employer that they could maintain an effective labor boycott. Faced with a strong and united front, the employer might concede to the strikers' demands. However, if the firm continued to operate during the strike, or seemed willing to bear the costs of a shutdown, strikers' confidence (and liquidity) would erode and more and more workers would cross the picket line. Faced with the possibility of permanent job loss, the remaining strikers would eventually return to work on the employer's terms.

Formally we consider a firm and a group of workers operating in a competitive labor market with market wage w_0. We assume that product-market power or firm-specific skills generate a "quasi rent" R per worker. In the absence of an effective union, the entire value of R is earned by the firm and workers earn the market wage. If a union is recognized we assume that bargaining results in a split of the quasi rent between workers and the firm. In this case, the wage is

$$w = w_0 + sR,$$

where $0 < s < 1$ represents a rent-splitting parameter. Treating s as fixed and normalizing the time horizon of the parties to 1, the "pie" in a strike over union recognition is therefore \$$sR$. If workers win the strike a wage premium is established, transferring \$$sR$ from profits to wages. If the strike fails, the wage remains at the market rate and the firm continues to earn all of the available quasi rent.

The other ingredient of our formal model is a specification of the delay costs of the parties. We assume that a strike of duration d imposes a cost $d \times c_w$ on workers and a cost $d \times c_f$ on the employer. Following Kennan and Wilson (1989) we assume that c_w and c_f are independent random variables whose realizations are observed asymmetrically: workers observe c_w but not c_f; the firm observes c_f but not c_w. Both parties know the distribution functions of their rival's costs and observe R.

The decision rules of the parties in this model depend only on the value of the prize *relative* to their delay costs. Let $v_w = sR/c_w$ and $v_f = sR/c_f$, and let $G_w(v_w)$ and $G_f(v_f)$ represent the induced distribution functions for v_w and v_f, respectively. The random variable v_j has the interpretation of the maximum profitable strike duration for party j (j = w, f). A party of "type" v_j is willing to endure a strike of length up to v_j rather than forgo the prize earned by the winner of the dispute.

[22] Erhlich (1974, p. 536). His conclusions are based on an analysis of editorials and news reports in the *National Labor Tribune* over the period from 1878 to 1885.

Equilibrium behavior of the parties in this bargaining game is described by a pair of concession functions which give the optimal "quitting times" associated with particular realizations of v_w and v_f (see Nalebuff and Riley 1985; and Fudenberg and Tirole 1986). In equilibrium, a firm with a higher realization of v_f (corresponding to a lower realization of delay costs) will hold out longer in anticipation of a capitulation by workers. Likewise, workers with a higher realization of v_w will hold out longer in anticipation of a firm capitulation. Depending on the distribution functions, workers and/or firms with sufficiently high delay costs may capitulate immediately (see Nalebuff and Riley 1985), implying no strike.

A difficulty with the war-of-attrition model is nonuniqueness. In general there is a continuum of pairs of equilibrium concession functions $\{T_w(v_w), T_f(v_f)\}$ with the property that if the firm follows the concession schedule T_f workers will follow T_w (and vice versa). This nonuniqueness is resolved if there is some finite probability that either party will strike forever (i.e., a positive probability of zero delay costs)—see Nalebuff and Riley (1985).[23]

Fudenberg and Tirole (1986) have derived general comparative statics results for a symmetric war-of-attrition model (i.e., a model with $G_w(v) = G_f(v)$). Assuming symmetry, they show that an increase in the payoff to the winner of the dispute leads to a longer delay time for each realization of costs. In our model this implies that an increase in quasi rent R leads to an increase in the expected capitulation time of both parties. Since we identify the wage increase conditional on a successful strike as a measure of R, factors that raise the wage settlement following a successful strike (but do not affect the delay costs of the parties) should lead to increases in the equilibrium capitulation times of both parties and an increase in expected strike duration.[24]

To derive comparative statics results for an asymmetric case, we assume that with probability $(1 - \pi_w)$ v_w is distributed uniformly on the interval $[0, R/\alpha]$ and that with probability π_w workers will never concede.[25] Similarly, we assume that with probability $(1 - \pi_f)$ v_f is distributed uniformly on the interval $[0, R/\beta]$ and that with probability π_f the firm will never concede. In this setup, α is a shift parameter for the distribution of workers' delay costs: an increase in α corresponds to a rightward shift in the distribution of workers' delay costs, leading to a leftward shift in the distribution of workers' maximum profitable delay times. Similarly, β is a shift

[23] Note that some strikes in our data set are in fact "infinite": in 5.4% of wage increase strikes, all the workers were replaced or the firm closed down.

[24] However, such factors should not affect the probability of a successful strike, since they raise the concession times of both parties.

[25] In terms of delay costs, this assumption implies that $c_w = 0$ with probability π_w, and that with probability $(1 - \pi_w)$ c_w is distributed on the range $[\alpha, \infty)$ with density αc_w^{-2}.

parameter for the distribution of the firm's delay costs: an increase in β corresponds to a rightward shift in the distribution of the firm's delay costs, leading to a leftward shift in the distribution of maximum profitable delay times for the employer.

In this example it can be shown that a decrease in α (or an increase in β) leads to an increase in the maximum strike duration for each quantile of v_w and a reduction in the maximum strike duration for each quantile of v_f. Thus, a downward shift in the distribution of delay costs of workers (or an upward shift in the distribution of delay costs of the firm) leads to an increase in the expected concession time of workers and a decrease in the expected concession time of the firm. Since the probability of a successful strike is just the probability that workers' capitulation time exceeds the firm's capitulation time, a decrease in workers' delay costs or an increase in firm's delay cost raises the probability of a worker success.

Variables that affect both the size of the quasi rent and the distributions of delay costs have potentially ambiguous effects on the expected capitulation times and the probability of a successful strike. In particular, if $R = R(X)$, $\alpha = \alpha(X)$, and $\beta = \beta(X)$, where X represents a set of characteristics of a particular dispute, then the equilibrium of the model depends on $R(X)/\alpha(X)$ and $R(X)/\beta(X)$. One particularly interesting case arises when the distribution of workers' delay costs is fixed across disputes but the distribution of firm's delay costs varies with the same factors that determine R. In this case, a variable that increases rents shifts out the distribution of maximum profitable strike durations for workers (since the size of the prize is higher) but has a smaller (or even negative) effect on the distribution of maximum profitable delay times for the firm (since although the prize is higher, delay costs are higher too). Such a variable will increase the expected capitulation time of workers, reduce the expected capitulation time of the employer, and increase the probability of a successful strike.

IV. An Empirical Analysis

A. Specification

Building on the descriptive evidence in Section II and the theoretical framework in Section III, we turn to a "structural" analysis of strike durations and outcomes. The building blocks of our empirical model are equations for the capitulation times of the two parties and an equation for the wage increase, conditional on a successful strike. As in the war-of-attrition model of animal conflict (Maynard Smith 1974), we assume that a strike ends when the strike duration exceeds the capitulation time of one of the parties. If workers concede first, the strike fails and wages return to their prestrike level. If the employer concedes first, the strike succeeds and a wage premium is established. The presence of compromise settlements poses a difficulty: we treat these as a third possible outcome with a separate specification for the maximum time until the parties will agree to a compromise.

Formally, we specify three equations for the latent random variables T_w, T_f, and T_c, representing the concession times of workers and the firm and the time until a compromise settlement, respectively. According to our theoretical model, T_w and T_f depend on the size of the available quasi rent (R), the parameters of the distributions of workers' and the firm's delay costs (α and β), and the actual realizations of the parties' delay costs. In principle, one could use a particular set of functional forms for the distributions of delay costs together with a set of assumptions on how observable and unobservable variables affect R, α, and β to derive functional forms for T_w and T_f. We follow an alternative "reduced-form" approach and specify a set of linear equations for the latent capitulation times:

$$T_w = X\beta_w + \varepsilon_w, \tag{1}$$

$$T_f = X\beta_f + \varepsilon_f, \tag{2}$$

and

$$T_c = X\beta_c + \varepsilon_c. \tag{3}$$

Here X is a vector of observed attributes shifting the equilibrium concession functions of the parties (industry and year effects, for example) and (ε_w, ε_f, ε_c) is a triple of random error terms, incorporating unobserved determinants of rents and delay costs and the specific realizations of delay costs. Observed strike duration is

$$T = \min[T_w, T_f, T_c]. \tag{4}$$

Equations (1)–(4) specify a competing-risks model with three (possibly correlated) risks (see Kalbfleish and Prentice 1980, chap. 7).

In addition to these equations, we specify an equation for the wage increase conditional on a strike success:

$$\Delta W = X\beta_d + \varepsilon_d. \tag{5}$$

We ignore wage outcomes in failed or partially successful strikes. As noted in Section II, the wage change conditional on a failed strike is almost always zero. Wage changes following partially successful strikes are typically nonzero (see fig. 1), and in principle we could add another equation for compromise wage settlements. In light of the small number of partially successful strikes, however, we have not done so.

In the estimates reported below, we assume that (ε_w, ε_f, ε_c, ε_d) have a joint normal distribution and that T_w, T_f, and T_c represent the logarithms of the latent concession times. The assumption of multivariate normality has a number of significant advantages. First, if ε_w, ε_f, and ε_c are assumed to be independent, the model reduces to three independent Tobit equations. Tobit-type estimates are a natural starting point for an analysis of censored duration data. Second, if compromise settlements are ignored, the normality assumption on ε_w and ε_f implies a simple probit model for the probability of a successful strike (see below). Third, joint normality allows us to incorporate arbitrary correlations between the unobserved determinants of the latent capitulation times. A model with correlated heterogeneity is especially attractive in light of the numerous unobservable variables that affect strike outcomes in our data. Finally, the assumption of joint normality allows us to model the correlations between unobserved determinants of strike duration and the unobserved determinants of the wage gain following a successful strike.

On the negative side, a normal competing-risks model imposes a restrictive functional form for the hazard rates of strike settlements (see Kalbfleish and Prentice 1980, pp. 24–25). An earlier version of this article contains an evaluation of the joint-normality assumption, including goodness-of-fit comparisons with a proportional hazards model with unrestricted baseline parameters (Card and Olson 1992). This analysis suggests that the normal competing-risks specification provides a reasonable fit to the data and successfully summarizes the effects of the observable variables on strike duration and the probability of a success.

B. Models of Strike Success and Wage Settlements

Before describing estimation results for the fully specified model, we present an initial analysis of the determinants of strike success and the wage increase conditional on a success. Abstracting from compromise settlements, equations (1) and (2) imply that the probability of a strike success is

$$\begin{aligned} P(\text{success}) &= P(X\beta_w + \varepsilon_w > X\beta_f + \varepsilon_f) \\ &= P[\varepsilon_f - \varepsilon_w < X(\beta_w - \beta_f)] \\ &= \Phi[X(\beta_w - \beta_f)/\sigma_1], \end{aligned}$$

where Φ is the normal distribution function and σ_1 is the standard deviation of $(\varepsilon_f - \varepsilon_w)$. This is a conventional probit model. The model of wage increases (eq. [5] above) is estimated without any attention to potential biases created by restricting the analysis to successful strikes. These biases are explored below in the discussion of the results for the full model.

The first two columns of table 4 present estimated coefficients from a probit model for the probability of a strike-success fit to the sample of strikes over wage increases. The third and fourth columns of the table present ordinary least squares (OLS) coefficient estimates of an equation

Table 4
Estimated Models for Probability of Successful Strike and Wage Increase following Successful Strike

	Probit Models for Probability of Success		OLS Models for Wage Increase in Successful Strikes	
	(1)	(2)	(3)	(4)
1. Ordered by labor organization	.49	.62	.019	.019
	(.11)	(.11)	(.008)	(.008)
2. Fraction of employees on strike	.34	.40	.039	.039
	(.16)	(.17)	(.011)	(.011)
3. Log number of strikers	−.06	−.06	−.007	−.007
	(.03)	(.03)	(.002)	(.002)
4. Fraction female employees	−.81	−.91	.007	.008
	(.33)	(.36)	(.022)	(.022)
5. Generic employees (Indicator)	.11	.08	−.002	−.002
	(.11)	(.12)	(.008)	(.008)
6. Strike in Massachusetts	−.07	.10	−.004	−.004
	(.19)	(.21)	(.012)	(.012)
7. Strike in Illinois	−.91	−.99	−.024	−.023
	(.20)	(.20)	(.017)	(.017)
8. Strike in Chicago	.87	.88	.018	.018
	(.22)	(.22)	(.017)	(.017)
9. Strike begun May 1–7, 1886	−.44	−.41	−.017	−.018
	(.19)	(.19)	(.014)	(.015)
10. Strike begun after May 7, 1886	−.21	−.19	−.015	−.015
	(.16)	(.17)	(.011)	(.011)
11. Strike duration (coefficient × 100)	. . .	−1.59	. . .	.003
		(.35)		(.031)
12. Strike duration squared (coefficient × 10,000)	. . .	.42	. . .	−.010
		(.17)		(.023)
13. R^2	. . .	. . .	.321	.322

NOTES.—See table 1 note for sources and definitions. All models include 11 industry and 5-year effects. Coefficients in cols. 1 and 2 are from probit model fit to sample of 1,026 strikes over wage increases. The average probability of a success is 0.519. Coefficients in cols. 3 and 4 are from ordinary least squares (OLS) regression fit to wage changes for subsample of 529 successful strikes. The mean wage increase (and standard deviation) are .136 and .072, respectively. Standard errors are in parentheses.

for the wage increase in the event of a successful strike. All of the models include a set of industry and year effects in addition to the covariates listed in the table.[26] In addition, for purely descriptive purposes, the models in

[26] The industries (and their relative frequencies in the sample) are tailors and clothing (5.8%), building trades (8.4%), food products (3.3%), wagon and carriage makers and similar machinery (2.8%), metal shops and implements (13.7%), mining (4.8%), shoes and boots (7.5%), textiles and shirts (3.0%), tobacco (13.7%), transportation (7.4%), wood products (6.6%), and miscellaneous industries (22.9%).

columns 2 and 4 include a quadratic function of observed strike duration. Within a war-of-attrition framework, the interpretation of these augmented models is problematic: we include them here to permit comparisons with other descriptive analyses of strike outcomes in these data.

The coefficients in table 4 confirm several findings of earlier studies and suggest a number of interesting hypotheses regarding the determinants of strike success and wage determination. First, as noted in the *Third Report*, strikes ordered by a labor organization were more likely to succeed.[27] Despite the importance attached to this effect by many early writers (see Adams 1905; and Moore 1911), its interpretation is unclear. On the one hand, the backing of a labor organization may affect strike costs—by providing organizational assistance to the strikers, for example, or by raising community support for the walkout.[28] On the other hand, union officers were often unwilling to sanction strikes with a low probability of success.[29] Some of the measured authorization effect surely reflects this selectivity.

The coefficients in rows 2 and 3 suggest that larger strikes were less likely to succeed, whereas strikes involving a larger fraction of the firm's employees were more likely to succeed. Again, there are a variety of interpretations of these effects. For example, an increase in the fraction of workers participating in a strike would be expected to lower the firm's chances of operating during a strike, thereby increasing its delay costs and raising the probability of success. Alternatively, a larger fraction of workers may have been willing to participate in strikes that were perceived as likely to succeed.

The employee composition effects (rows 4 and 5) suggest that strikes involving female workers were significantly less likely to succeed, whereas strikes involving generic employee groups had about the same success rate as other strikes. The geographic variables show similar success rates in New York, Massachusetts, and Chicago, but much lower rates for strikes

[27] Unlike us, Friedman (1988, table 4) finds that authorized strikes were no more likely to succeed than other strikes before 1887, although his sample includes strikes for all causes. When we expand our sample to include all strikes, we still find a significant positive effect of union authorization. We have also estimated models that interact union authorization with 1885 and 1886 dummies to capture any differences due to the rise in the Knights of Labor in those years. These interactions are positive but insignificant, showing no less effectiveness of union-ordered strikes in these years than before.

[28] Boycotts were used by trade unions and especially by the Knights of Labor in the 1880s to increase pressure on employers during strikes. See Foner (1975, pp. 48–50) for a series of specific examples.

[29] Janes (1916) and Ulman (1955) describe the mechanisms put in place by national unions in the 1880s and 1890s to prevent local union leaders and/or members from engaging in strikes. Adams (1905, p. 181) argued that officers of national unions had "much more to lose in place, power, and prestige, by an unsuccessful strike" than local leaders or members.

in other parts of Illinois.[30] Finally, the coefficients in rows 9 and 10 show a sharp decrease in success rates for strikes launched during and after the "general strike" in May 1886 (see Sec. *E* below). We have also estimated models with seasonal dummy variables for the starting date of the strike and models that include the prestrike wage of strikers. In neither case is the estimated effect large or statistically significant.

The addition of a quadratic function of strike duration to the probit model (col. 2) confirms the conclusion from table 3: workers were much less likely to win long strikes. The coefficient estimates imply a 40-percentage-point reduction in the probability of a successful strike after a 100-day stoppage. We stress, however, that this is a descriptive correlation rather than a causal effect between duration and the strike outcome. In our model, the winner of the strike and strike duration are jointly determined by the capitulation hazards of the two parties. The fact that longer strikes are more likely to be won by the firm could be because of true duration dependence in the capitulation hazards of the parties, or it may simply reflect unobserved heterogeneity in relative capitulation times. For example, the data could be generated by two types of bargaining pairs: the first bargaining pair type composed of firms with a "short" capitulation time and workers with a slightly longer capitulation time, and a second bargaining pair type, where workers have a "long" capitulation time and firms have a "slightly longer" capitulation time.

A comparison of the coefficients for the wage increase models (cols. 3 and 4) with the coefficients of the probit models reveals an interesting regularity: variables that raise (lower) the likelihood of a successful strike also raise (lower) the wage conditional on workers winning the dispute. Indeed, the correlation of the 10 coefficient estimates in column 1 with the corresponding estimates in column 3 is 0.67. The same pattern and degree of correlation are revealed by the industry effects—industries with a higher probability of a strike success also have larger wage gains conditional on a success (see below). The strike duration coefficients in columns 2 and 4 are an exception to this rule. Whereas the probability of successful settlement declines with strike duration, the wage increase conditional on workers' winning the strike does not. The invariance of the wage settlement to the duration of the strike lends further credence to the view that wins and losses were discrete outcomes in the labor conflicts of the 1880s.

What interpretation does our theoretical model offer for the finding that the probability of strike success varies with the size of the potential wage gain if the strike succeeds? As noted above, an increase in the size of the prize should not necessarily raise the equilibrium win rate of one party or the other: if their delay costs are constant, both parties will be willing to

[30] We found no significant differences in strike success rates or in the wage increases for successful strikes in New York City or Boston.

hold out longer for a larger prize. The systematic correlation between the coefficient estimates in columns 1 and 3 suggests a different hypothesis—that as the potential rents are increased, the distribution of employer's delay costs also shifts upward, leading to a smaller increase in the net payoff from winning the strike. If potential rents and firm's delay costs are systematically correlated, a war-of-attrition model predicts higher equilibrium win rates for employees in strike situations involving greater rents. Alternative models of strike durations and wage outcomes may lead to the same prediction.

C. Competing Risk Models for Strike Duration

Table 5 reports estimation results for several versions of the strike duration model composed of equations (1)–(4). Column 1 of the table presents a simple linear regression model for the logarithm of strike duration. Columns 2–4 present estimates from independent Tobit models fit to successful, failed, and partially successful strikes, respectively. Finally, columns 5–7 present estimates from a three-equation competing-risks model, allowing unrestricted correlations between the residual components of the three latent durations.

If the distinction between strike outcomes is uninformative, the three concession time equations share the same coefficient vector and the competing-risks model degenerates to a single equation for log strike duration. Comparisons of the coefficients in column 1 with the outcome-specific coefficient estimates suggest that this restriction is rejected. For example, the union authorization variable (row 1) has a much larger effect on workers' capitulation time than on the firm's (cf. cols. 2 and 3 or 5 and 6). By raising the capitulation time of workers *relative* to the employer, union authorization is predicted to increase the likelihood of a successful strike. This inference is confirmed by the probit coefficients in table 4.

Analogous differences emerge in the effects of the other covariates. For example, an increase in the fraction of employees involved in the strike raises the capitulation time of workers and lowers the capitulation time of employers, implying a net positive effect on the probability of a successful strike. Again, this conclusion is confirmed by the probit coefficients in table 4.

Comparing the independent Tobit models and the joint competing-risks model, we find that most of the coefficient estimates are similar, although the joint model often leads to a bigger difference between the coefficients of the worker and firm capitulation time equations. The coefficients that change most dramatically between the two specifications are those involving female employees. An increase in the fraction female appears to reduce the union's and increase the firm's capitulation time. The estimated correlations of the error terms are shown in row 12 of table 5. These correlations are all positive, although only the correlation

Table 5
Estimated Models for Time to Capitulation or Compromise

	OLS Log Duration Model (1)	Independent Tobit Models of Capitulation Times			Competing Risks Correlated Tobit Model		
		Workers (2)	Firm (3)	Compromise (4)	Workers (5)	Firm (6)	Compromise (7)
1. Ordered by labor organization	.60	1.01	.16	.93	.90	.30	.95
	(.10)	(.12)	(.12)	(.18)	(.13)	(.16)	(.23)
2. Fraction of employees on strike	−.18	.15	−.39	−.47	.18	−.37	1.24
	(.15)	(.18)	(.18)	(.33)	(.17)	(.21)	(.84)
3. Log number of strikers	.16	.18	.11	.09	.13	.19	.06
	(.03)	(.04)	(.03)	(.06)	(.03)	(.04)	(.07)
4. Fraction female employees	.13	.07	.46	−.36	−.41	.79	1.24
	(.28)	(.34)	(.35)	(.53)	(.38)	(.44)	(.84)
5. Generic employees (indicator)	−.04	.04	.06	−.36	−.03	−.08	−.21
	(.10)	(.13)	(.13)	(.20)	(.11)	(.14)	(.23)
6. Strike in Massachusetts	.32	.06	.35	−.17	.25	.34	.51
	(.17)	(.24)	(.20)	(.36)	(.19)	(.22)	(.41)
7. Strike in Illinois	−.03	−.52	.43	−.43	−.50	.65	−.35
	(.16)	(.18)	(.25)	(.26)	(.21)	(.29)	(.35)
8. Strike in Chicago	.15	.50	−.27	.38	.48	−.43	.55
	(.18)	(.21)	(.26)	(.30)	(.21)	(.30)	(.37)
9. Strike begun May 1–7, 1886	.63	.36	.67	−.01	.44	.95	.06
	(.16)	(.20)	(.23)	(.27)	(.24)	(.28)	(.32)
10. Strike begun after May 7, 1886	.38	.31	.31	−.15	.20	.47	.18
	(.15)	(.19)	(.18)	(.24)	(.17)	(.20)	(.28)
11. σ	1.17	.94	1.07	.90	1.06	1.34	1.37
	(.06)	(.03)	(.03)	(.05)	(.04)	(.08)	(.27)
12. Estimated correlation parameters (correlated Tobit model):							
Correlation of workers and firm times (ρ_{wf})						.66	
						(.27)	
Correlation of workers and compromise times (ρ_{wc})						.52	
						(.51)	
Correlation of firm and compromise times (ρ_{fc})						.17	
						(.40)	

NOTE.—Models are fit to a sample of 971 strikes over wage increases for which duration information is available. See the text. All models include 11 industry and 5-year effects. Standard errors are in parentheses.

between ε_w and ε_f is statistically significant. In the context of our model, the positive correlation of ε_w and ε_f may reflect unobserved differences in the size of the quasi rents across bargaining pairs. Such differences would be expected to generate a positive correlation between the capitulation times of the parties.

While many of the observed strike characteristics have significant effects on the times until a success or failure, only the authorization variable has a significant effect on the time until a compromise. The year and industry effects in the compromise equation are also poorly determined. In part, this may be attributable to the small number of compromise cases. Two variants of the model in table 5 were estimated to gain some further insight into the nature of compromises. In the first case, we constrained the parameters of the compromise equation (β_c) to equal the coefficients of the firm capitulation equation. In the second case, we constrained β_c to equal the coefficient vector in the employees' capitulation equation. Neither of these restrictions fits the data. Compromises thus appear to represent a different outcome than either successful or failed strikes.[31]

D. A Joint Model of Strike Duration and Wage Outcomes

The three-equation competing-risks model of strike duration can be extended to a "complete" model of strike outcomes by adding an equation for the wage increase conditional on strike success. A difficulty with this model is computational complexity: the system of equations (1)–(5) has over 100 parameters if we include unrestricted industry and year effects in all the equations. Since the estimated correlations of the compromise equation with the worker and firm capitulation equations are insignificant (see table 5), a reasonable strategy is to drop the equation for partially successful strikes and treat these as independently censored observations. Following this approach, it is possible to estimate the coefficients of the employer and employee capitulation time equations (β_w, β_f), the coefficients of the wage-increase equation (β_d), and the correlation matrix of (ε_w, ε_f, ε_d) using the entire sample of wage-increase strikes.

Coefficient estimates for this extended model are presented in table 6. Columns 1 and 2 of the table present the worker and firm capitulation equations, respectively, while column 4 presents the wage-increase equation. Column 3 reports the *difference* in the coefficient estimates for the worker and firm equations, that is, estimates of ($\beta_w - \beta_f$). As noted earlier, the coefficients in a probit model for the likelihood of a successful strike

[31] Compromise settlements tend to occur after relatively long strikes. The median time to a compromise in the data is 12 days, compared with 6 days for a successful strike and 11 days for a failed strike.

Table 6
Estimated Models for Time to Capitulation and Wage Increase Given Successful Strike

	Capitulation Time Equations			Wage Increase Equation
	Workers (β_w) (1)	Firm (β_f) (2)	$\beta_w - \beta_f$ (3)	(4)
1. Ordered by labor organization	.81	.34	.47	.036
	(.13)	(.14)	(.22)	(.008)
2. Fraction of employees on strike	.18	−.31	.49	.051
	(.18)	(.18)	(.23)	(.011)
3. Log number of strikers	.13	.19	−.06	−.008
	(.03)	(.03)	(.01)	(.002)
4. Fraction female employees	−.41	.59	−1.00	−.035
	(.32)	(.38)	(.49)	(.024)
5. Generic employees (indicator)	−.01	−.07	.06	.002
	(.11)	(.12)	(.17)	(.008)
6. Strike in Massachusetts	.24	.34	−.10	−.005
	(.19)	(.21)	(.26)	(.013)
7. Strike in Illinois	−.35	.49	−.84	−.067
	(.20)	(.26)	(.30)	(.018)
8. Strike in Chicago	.33	−.32	.65	.059
	(.20)	(.27)	(.35)	(.019)
9. Strike begun May 1–7, 1886	.49	.94	−.45	−.030
	(.19)	(.21)	(.23)	(.015)
10. Strike begun after May 7, 1886	.22	.47	−.25	−.019
	(.17)	(.18)	(.23)	(.012)
11. σ	1.122	1.310	...	.069
	(.033)	(.048)		(.004)
12. Estimated correlation parameters:				
ρ_{wd}		.30		
		(.14)		
ρ_{fd}		−.23		
		(.05)		
ρ_{wf}		.89		
		(.23)		

NOTE.—Models are fit to a sample of 971 strikes over wage increases for which duration information is available. See the text. All models include 11 industry and five year effects. Standard errors are in parentheses.

are proportional to ($\beta_w - \beta_f$) (ignoring the presence of compromise settlements). Thus, an informal specification check of the extended model is obtained by comparing the estimates in column 3 of table 6 to the simple probit coefficients in column 1 of table 4. By the same token, the wage coefficients in column 4 of table 6 can be compared to the OLS coefficients in column 3 of table 4.

The estimated coefficients of the capitulation time equations are very similar to the estimates obtained in table 5. Evidently, the treatment of partially successful strikes makes little difference to the estimates of duration models for successful and failed strikes. The estimated wage change coefficients in table 6 are also similar to the OLS estimates in table 4, although uniformly larger in magnitude, as would be expected if the OLS estimates are attenuated by selection bias. Finally, the estimates of $(\beta_w - \beta_f)$ in column 3 are very similar to the probit coefficients in table 4.[32]

A comparison of the estimates in column 3 with the estimates in column 4 confirms that strike characteristics that increase the *relative* capitulation time of workers also raise the wage increase following a success. The relation between relative capitulation times and wage gains extends to the unobserved determinants of strike duration and wage increases. Unobserved components of workers' capitulation time tend to raise the wage settlement ($\rho_{wd} > 0$ in row 12), whereas unobserved components of the firms' capitulation time tend to lower the wage settlement ($\rho_{fd} < 0$). As a consequence, the unobserved determinants of strike success are positively correlated with the unobserved determinants of the wage settlement.[33]

The parallelism between relative capitulation times and wage gains also applies to the pattern of the industry effects. This is illustrated in figure 2, which plots the industry effects from the wage-settlement equation against the corresponding *differences* in the industry effects of the worker and firm capitulation equations. Each point in the figure corresponds to a different industry. Apart from the tobacco industry, the points lie on a positively sloped line, confirming the strong link between workers' relative ability to withstand a strike and their expected wage gain conditional on a success. It is important to note that the correlation across industries between the average wage increase for a successful strike is correlated with the difference in capitulation times, and not average strike duration.[34] In fact, there is no significant interindustry relationship between the average duration of strikes and the average industry wage effect conditional on a successful strike.

The tobacco industry is an outlier: although tobacco strikes had high success rates, wage increases conditional on a success were below average. There are several possible explanations for this finding. Technological changes in the tobacco industry during the 1880s led to the gradual replacement of highly skilled cigar rollers by less skilled cigar molding operatives (see Ware 1929, chap. 11). Pressure from this ongoing "deskilling" may account for the relatively modest wage increases in our sample. Another factor was the relatively high level of union organization in the industry. Unlike most other workers engaged in wage increase strikes, tobacco workers may have *already* earned substantial union wage premiums. Finally, the bitter rivalry between two different cigarmakers unions

[32] Note that the probit coefficients are estimates of $(\beta_w - \beta_f)/\sigma_1$, where σ_1 is the standard deviation of $(\varepsilon_f - \varepsilon_w)$. However, the estimates in rows 11 and 12 of table 6 imply that σ_1 is very close to one.

[33] The unobserved component of the probability of strike success is $(\varepsilon_w - \varepsilon_f)$, which is positively correlated with ε_d given $\rho_{wd} > 0$ and $\rho_{fd} < 0$.

[34] Similar effects were found using a simple Tobit model that treats compromise settlements as censored.

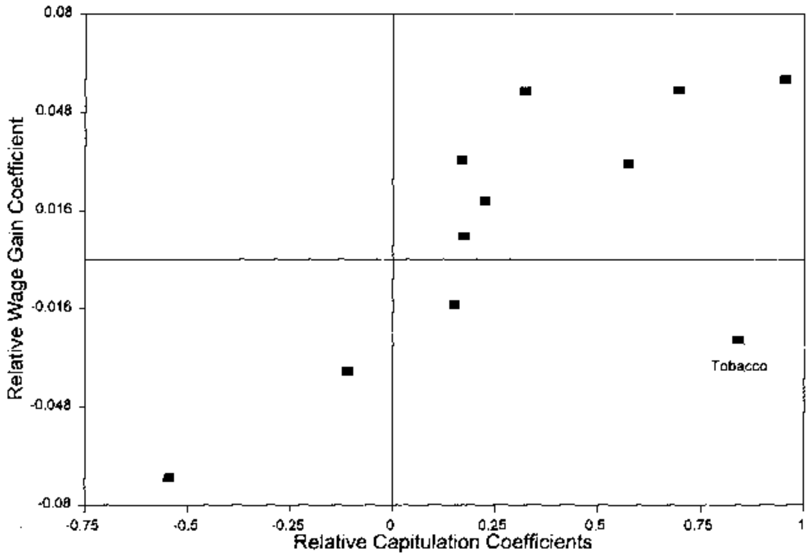

FIG. 2.—Wage gain and relative capitulation coefficients, by industry

in New York City may have contributed to the unusual character of strikes and wage settlements in the industry (Ware 1929).

E. Effects of the 8-Hour Movement and the Haymarket Affair

The coefficient estimates in tables 4–6 all point to significant changes in strike outcomes during and after the first week of May 1886. Both worker and firm capitulation times increased dramatically in the wave of strikes from May 1 to May 7, and continued above the level of earlier strikes throughout the remainder of the year.[35] The increase was significantly greater for firms, implying a 15-percentage-point reduction in the probability of a successful strike for disputes beginning in the first week of May and an 8% lower success rate for strikes in the latter part of the year. These changes were accompanied by reductions in the size of wage increases conditional on a strike success: 3% lower for strikes in the first week of May and 2% lower for later strikes.

Within the framework of our theoretical model, these estimates suggest that the wave of strike activity in May 1886 involved groups of workers with relatively low potential gains from striking (controlling for other observable characteristics). Even in cases where the strikes succeeded, workers were able to achieve only modest wage premiums. The shift in

[35] The coefficients are normalized relative to strikes beginning in January–April of 1886. The year effects show that strike durations and wage outcomes were fairly stable between 1884 and early 1886.

composition toward employees with low potential gains from striking was associated with a sharp reduction in the probability of success.

Nevertheless, it is interesting to note that in terms of *observable* characteristics, the groups involved in strikes during May 1886 were not too different from strikers in earlier periods. To track the composition of strikes we fit a probit model (identical to the probit in col. 1 of table 4) to strike outcome data from 1881–85. We then used this model to compute the predicted success rate of wage increase strikes in 1886, and compare the predicted and actual rates. The results show surprisingly little change in the predicted probability of strike success from 1885 to 1886.[36] Although strikes in 1886 were larger and were more likely to involve generic employee groups than earlier strikes, these differences can account for only a small fraction of the observed decline in success rates.

V. Summary and Conclusions

Our analysis of strikes in the early 1880s leads to two main conclusions. First, strike outcomes were fundamentally discrete. For strikes over wage increases, a failed strike meant a return to work at the prestrike wage. A successful strike, on the other hand, meant a significant wage increase (averaging 13%). Ninety percent of strikes were resolved by one of these two outcomes. Second, win/loss probabilities were proportional to the size of the wage gain if the strike succeeded. Analyzing patterns across larger and smaller strikes, strikes with higher and lower participation rates, strikes in different industries, and strikes before, during, and after the wave of unrest in May 1886, we find a consistent pattern linking the wage premium for a successful strike to the probability of success.

The discreteness of strike outcomes and other qualitative evidence lead us to interpret disputes over wage increases as contests to determine the bargaining status of workers. If a strike succeeded, the strikers' bargaining power was recognized and a wage premium—the equivalent of a union wage gap—was established. Otherwise, employers continued to earn all the potential rents. This interpretation maps into a war-of-attrition model in which the prize for the winner of the dispute is a share of the rents. The theoretical model highlights the importance of the parties' strike costs *relative to* the size of the available rents. Within this framework, the proportionality between win/loss rates and the wage premium for a successful strike can be interpreted as evidence that employers with greater rents had higher delay costs during a work stoppage.

[36] This is also true of strikes over other issues. Using an index based on the probability of winning a wage increase strike, the observed characteristics of hours-related and other strikes in May 1886 account for very little of the decline in success rates in that period.

Regardless of theoretical interpretation, the finding that strikes were most likely to succeed in situations where workers had the largest potential wage gains from collective action provides an interesting perspective on union organizing policies after 1886. With the rise of the American Federation of Labor, trade unions in the United States moved toward narrowly focused craft unionism. Our results suggest that this policy would ensure the greatest likelihood of strike success in the difficult period of the 1890s.

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